

UNIVERSITY OF
ILLINOIS LIBRARY
AT URBANA-CHAMPAIGN
BOOKSTACKS

B24E 2-3

THE HECKMAN BINDERY, INC. North Manchester, Indiana				BINDING COPY				
H or V	JUST	FONT	SLOT	TITLE	PERIODICAL <input type="checkbox"/> CUSTOM <input type="checkbox"/> STANDARD <input type="checkbox"/> ECONOMY <input type="checkbox"/> THESIS <input type="checkbox"/>	BOOK <input type="checkbox"/> CUSTOM <input type="checkbox"/> MUSIC <input type="checkbox"/> ECONOMY <input type="checkbox"/>	NO VOLS THIS TITLE	LEAD ATTACH
H	CC	1W	22		ACCOUNT	LIBRARY NEW	RUB OR SAMPLE	TITLE I.D.
			21	PTY	6672 001		6832	WKE 488
			20	WORKING	ACCOUNT NAME			
			19	PAPER	UNIV OF ILLINOIS			
H	CC	1W	8	1983-84	ACCOUNT INTERNAL I.D.		ISSN	
			7	NO.991-1009	B01912400			
H	CC	1W		330	I.D. #2	NOTES	BINDING FREQUENCY	WHEEL
				B385<"CV">	STX4			SYS I.D.
				no.991-1009	COLLATING			39276
				COB. 2	35			
H	CC	7W		KIMPRINT>	ADDITIONAL INSTRUCTIONS			
				U. of ILL.	Depreciated Lot=201 Items=121 (100)=12V			
				LIBRARY	1CR2ST3CR MARK ST # B4 911			
				UREANA				
SEP. SHEETS					PTS. BD. PAPER		TAPE STUBS	
CLOTH EXT.					FILLER		GUM	
POCKETS					SPECIAL PREP		LEAF ATTACH	
PAPER					BUCK		CLOTH	
INSERT MAT					ACCOUNT LOT NO		JOB NO	
PRODUCT TYPE					ACCOUNT PIECE NO.		PIECE NO	
HEIGHT					GROUP CARD		VOL THIS TITLE	
COVER SIZE					X		001247943	

85
994
P2



BEBR

FACULTY WORKING
PAPER NO. 994

The Present Value Model of Stock Prices: Empirical
Tests Based on Instrumental Variables Estimators

Louis O. Scott

College of Commerce and Business Administration
Bureau of Economic and Business Research
University of Illinois, Urbana-Champaign

BEBR

FACULTY WORKING PAPER NO. 994


College of Commerce and Business Administration

University of Illinois at Urbana-Champaign

November 1983

The Present Value Model of Stock Prices:
Empirical Tests Based on Instrumental Variables Estimators

Louis O. Scott, Asst. Professor
Department of Finance



Digitized by the Internet Archive
in 2011 with funding from
University of Illinois Urbana-Champaign

<http://www.archive.org/details/presentvaluemode994scot>

ABSTRACT

The variance bounds tests of the present value model of stock prices are re-examined in this paper. A direct test of the model based on generalized instrumental variables estimation is proposed as an alternative and it is shown that this approach has several advantages over the variance bounds tests. The test based on instrumental variables estimation is modified to handle the case in which the percentage changes in real dividends and real stock prices are stationary processes. The tests are applied to quarterly data for the Standard & Poor's Index of 500 Common Stocks and the results are much more conclusive than those obtained by the previous tests.

THE PRESENT VALUE MODEL OF STOCK PRICES:
EMPIRICAL TESTS BASED ON INSTRUMENTAL VARIABLES ESTIMATORS

Shiller (1981a and 1981b) and LeRoy and Porter (1981) have tested the present value model of stock prices by examining the implicit restrictions on the variation of stock prices. Their results suggest that actual stock prices vary too much to be consistent with this model. If we examine their results closely, we find that Shiller does not construct formal statistical tests of the model and that most of the tests in LeRoy and Porter are not statistically significant because the standard errors of the variance estimates are quite large. Some critics have argued that the relevant time series are not covariance stationary, even after the removal of a time trend, and that the variance estimates are therefore unreliable. An alternative is to assume that the percentage changes in dividends, earnings, and stock prices are covariance stationary time series. Shiller (1981a) considers this alternative assumption, but notes that it does not lead to tractable variance bounds for stock prices. In this paper, a test of the present value model based on an instrumental variables estimator is developed as an alternative to the variance bounds test, and the test is extended to handle the case in which the percentage changes in dividends, earnings, and stock prices are covariance stationary. This latter specification requires the use of Hansen's (1982) generalized method of moments estimator.

I. Variance Bounds in the Present Value Model of Stock Prices

The present value model of stock prices has the following form:

$$P_t = E_t \left\{ \sum_{j=1}^{\infty} \beta^j D_{t+j} \right\},$$

where P_t is the asset price at the end of period t and D_t is the dividend paid during period t . E_t is the conditional expectations operator, conditional on information available in period t . E without a subscript will represent the unconditional expectations operator. P_t and D_t are expressed in real terms. This simple expectations model follows from the expected real rate of return being constant and is generally associated with risk neutrality in asset pricing models.¹ If we let

$P_t^* \equiv \sum_{j=1}^{\infty} \beta^j D_{t+j}$, we observe that the model implies

$$(1) \quad P_t^* = P_t + \eta_t,$$

where η_t is the forecast error. The assumption of rational expectations (or market efficiency) requires that P_t and η_t be uncorrelated, so that we get

$$\text{Var}(P_t^*) = \text{Var}(P_t) + \text{Var}(\eta_t),$$

and we conclude that $\text{Var}(P_t) \leq \text{Var}(P_t^*)$. Thus, we have an upper bound on the variation of stock prices.

Two problems are encountered when one attempts a test of this bound on the variation of stock prices. First, the time series must be covariance stationary; that is, the variances and covariances of the series must be finite and must not depend on time. Shiller removes a long-term trend from his series on dividends and prices and applies the variance restrictions to the detrended series. LeRoy and Porter argue that there are no apparent trends in their adjusted series for earnings and prices and that further adjustments are not necessary. The second problem

deals with the estimation of the variance of P_t^* (or a detrended P_t^*), which is not observable. Shiller calculates this series recursively by assuming a value at the end of the sample period and then computes the sample variance, but he does not develop a statistical test of the model. LeRoy and Porter formulate and estimate a finite-parameter bivariate time series model to compute the variances. This method requires the researcher to formulate a time series model and then test the variance bounds conditional on the formulated model. Singleton (1980) has shown that the variance of a series like P_t^* can be estimated in the frequency domain if we have a value for the discount factor.² The hypothesis tests that a researcher might develop for these variance estimators will depend on the large sample distribution theory for variance estimators. It is well-known that these large sample distributions depend on the effects of fourth cumulants of the innovations of the process generating the observed time series, and for this reason the distributions typically used for hypothesis testing are not robust if we drop the assumption that the innovations are multivariate normal.³ It is extremely difficult to estimate the standard errors of the variance estimates if we want to relax the assumption of normally distributed innovations (or the assumption that the fourth cumulants are all zero). The large sample distributions for many of the standard econometric estimators (such as ordinary least squares, generalized least squares, or instrumental variables) do not require an assumption that the innovations or error terms are normally distributed. For this reason, the tests based on instrumental variables estimators developed in the next section impose a less restrictive set of assumptions.

LeRoy and Porter developed several tests for the implied variance bounds. Even though their point estimates indicate rejection of the present value model, the confidence intervals are so large that the model is rejected in only a few of the cases examined. In a chapter of my dissertation (1982, Ch. 5), I calculated frequency domain variance estimators and found the point estimates for the variances of the price series to be many times greater than the estimates for the upper bounds, but the standard errors associated with these estimates are so large that the tests of the model are barely significant at the 5% level. It appears that these variances are not being estimated with much precision. More recently, Flavin (1982) has questioned the finite sample properties of the frequency domain variance estimators. One can hardly argue that these tests of the present value model are conclusive.

II. An Alternative Test Based on Generalized Instrumental Variables Estimation

In an unpublished paper, Geweke (1979) has shown that regression tests of the simple expectations models can be more powerful than the variance bounds tests. This testing procedure has been the one most frequently applied in other tests of simple expectations models (e.g., tests of forward rates as unbiased predictors of spot rates in foreign exchange markets). The regression tests, however, cannot be applied directly to the present value model of stock prices. If the series P_t^* were observable, then the relationship in equation (1) could be easily tested by a least squares regression of P_t^* on a constant and P_t . To derive a regression model, we need to restate equation (1) in

terms of observable dividends and stock prices. First rewrite the equation under the alternative hypothesis using the lag operator (L : $Lx_t = x_{t-1}$):

$$\frac{\beta L^{-1}}{1-\beta L^{-1}} D_t = a + bP_t + \eta_t.$$

Now we multiply both sides of the equation by the filter $\frac{1-\beta L^{-1}}{\beta L^{-1}}$ to get

$$(2) \quad D_t = a - bP_t + \frac{b}{\beta} P_{t-1} + u_t,$$

where $u_t = (\frac{1}{\beta} L - 1)\eta_t$. We now have an equation that is a function of observable time series and an error term, but the equation does not satisfy all the assumptions of the classical regression model. The forecast errors η_t represent forecast errors for future dividends and are serially correlated. The error term u_t , which is a function of the forecast errors, will be serially correlated in most cases. The more important problem is that P_t and u_t are correlated; therefore, we must use an instrumental variables (IV) estimator for the parameters of the equation. Under the null hypothesis of the present value model, $a=0$, $b=1$, and $E_t(\eta_{t+j})=0$ for $j \geq 0$. Under the alternative hypothesis, we relax the restrictions on a and b and preserve the restriction on the forecast errors. Because the error term of equation (2) is a function of η_t and η_{t-1} ,

$$E(D_{t-j} u_t) = E(P_{t-j} u_t) = 0 \quad \text{for } j > 0,$$

but $E(P_t u_t) \neq 0$. In addition, u_t should be uncorrelated with any variable (e.g., earnings) dated $t-1$ or earlier. A natural set of instruments for this problem should include lagged dividends, the lagged

stock price, and lagged values of variables which are important for predicting dividends. Lintner (1956) and Fama and Babiak (1968) have shown that earnings play an important role in the behavior of dividends over time; therefore lagged earnings should be included.

The relevant time series must have finite second moments in order to estimate the parameters in equation (2), and this condition is normally satisfied by requiring that the time series be stationary, or at least covariance stationary.⁴ This condition is necessary for estimating the variances in the variance bounds tests and it is equally important for estimating the parameters in equation (2). One approach is to follow Shiller and remove a long-term trend from the data as follows: $p_t = P_t(1+g)^{-t}$ and $d_t = D_t(1+g)^{-t}$. The discount factor for the model using detrended data becomes $\gamma = \beta(1+g)$, and the following equation can be derived for the detrended series:

$$(3) \quad d_t = a - b p_t + \frac{b}{\gamma} p_{t-1} + u_t.$$

For equation (3), we still have the condition that $E_{t-1}(u_t) = 0$ and we must use an IV estimator. To motivate the alternative hypothesis in which $a \neq 0$ and $b \neq 1$, we can imagine a world in which the stock market correctly discounts the constant and the long-term trend of dividends, but systematically overestimates or underestimates the deviations of dividends about the long-term trend. If the market overestimates these deviations from the long-term trend, then the coefficient b should have a value less than one indicating that stock prices vary too much. The estimation problem for equation (3) does simplify because it can be shown that the error term for this specification is white noise. This

result is shown in an appendix for the case in which d_t has a linear multiple time series representation. Because the error term is not serially correlated, we can directly apply Amemiya's (1974) nonlinear two-stage least squares estimator.

In the alternative specification, I assume that the dividend process is not mean-reverting. For this second case, I assume that the percentage change in dividends and stock prices as well as the price-dividend ratio $(\frac{\Delta D}{D}, \frac{\Delta P}{P}, \frac{P}{D})$ are stationary. Shiller (1981a) notes that the terminal condition for the present value model is not necessarily satisfied if we assume that only $\frac{\Delta D}{D}$ and $\frac{\Delta P}{P}$ are stationary. If the terminal condition is not satisfied, we have the undesirable result that there is no solution for the stock price. Shiller imposes the terminal condition by requiring the price-dividend ratio to be stationary.⁵ The relationship in equation (1) is now modified as follows:

$$\frac{P_t^*}{D_t} = a^* + b^* \frac{P_t}{D_t} + \frac{\eta_t}{D_t},$$

and again $a^*=0$ and $b^*=1$ for the present value model. P_t^* is again a function of future dividends and we apply the same filter as before to both sides of the equation after rewriting it as follows:

$$P_t^* \equiv \frac{\beta L^{-1}}{1-\beta L^{-1}} D_t = a^* D_t + b^* P_t + \eta_t$$

$$D_t = a^* \left(\frac{1}{\beta} D_{t-1} - D_t \right) + b^* \left(\frac{1}{\beta} P_{t-1} - P_t \right) + \left(\frac{1}{\beta} \eta_{t-1} - \eta_t \right)$$

$$D_t = a D_{t-1} - b P_t + \frac{b}{\beta} P_{t-1} + \frac{-\eta_t + \frac{1}{\beta} \eta_{t-1}}{1+a^*},$$

where $a = \frac{a^*}{\beta(1+a^*)}$ and $b = \frac{b^*}{1+a^*}$. Then equation can then be rewritten as a function of stationary time series plus an error term.

$$(4) \quad \frac{D_t}{D_{t-1}} = a - b\left(\frac{P_t}{D_{t-1}}\right) + \frac{b}{\beta}\left(\frac{P_{t-1}}{D_{t-1}}\right) + u_t,$$

where $u_t = \frac{-\eta_t + \frac{1}{\beta}\eta_{t-1}}{(1+a^*)D_{t-1}}$. Under the null hypothesis for the present value model, $a=0$ and $b=1$. The error u_t again has the property that it is uncorrelated with variables dated $t-1$ and earlier, but it is not possible to show that it is serially uncorrelated. Note that $E(u_t(\frac{P_t}{D_{t-1}})) \neq 0$, and again we must use an IV estimator. Because the error term may be serially correlated, we must use Hansen's (1982) generalized method of moments (GMM) estimator to estimate the parameters of equation (4). It is convenient to think of the GMM estimator in this case as an IV estimator generalized to allow for a serially correlated error term. The exact estimator used is presented in the next section.

The IV estimators for both cases (equations (3) and (4)) can be used to construct χ^2 tests of the restrictions for the present value model (a joint test that $a=0$ and $b=1$). This test has several advantages over the variance bounds test. The variance bounds test is based on an inequality restriction which may include a substantial amount of slack, whereas the test based on the IV estimators constitutes a direct test of the model. The IV estimation test has three additional properties: (1) an initial estimate of the discount factor is not required, (2) we do not need to assume that the fourth cumulants of the innovations are zero nor do we need to assume normality, and (3) we do not need to formulate a finite parameter time series model for dividends and stock prices.

III. The IV and GMM Estimators

To develop the IV estimator for equation (3), first rewrite the equation in vector notation as follows:

$$\underline{u} = \underline{d} - a\underline{1} + b\underline{p} - \frac{b}{\gamma} \underline{p}_{-1},$$

where the vectors are of length T, the sample size, and $\underline{1}$ is a vector of ones. \underline{p}_{-1} is the detrended price lagged one period. For the instrumental variables, we use a constant, d_{t-1} , p_{t-1} , and x_{t-1} (detrended earnings lagged one period), so that the matrix of instrumental variables is

$$Z \equiv [\underline{1} \quad \underline{d}_{-1} \quad \underline{p}_{-1} \quad \underline{x}_{-1}].$$

Let $\underline{\theta}' = (a, b, \gamma)$ and we estimate $\underline{\theta}$ by⁶

$$\min_{\underline{\theta}} \ell = \underline{u}' Z (Z' Z)^{-1} Z' \underline{u}.$$

The covariance matrix for $\underline{\theta}$ is computed as follows:

$$\text{Var}(\hat{\underline{\theta}}) = \hat{\sigma}_u^2 [(\frac{\partial \underline{u}}{\partial \underline{\theta}})' Z (Z' Z)^{-1} Z' (\frac{\partial \underline{u}}{\partial \underline{\theta}})]^{-1},$$

where $\frac{\partial \underline{u}}{\partial \underline{\theta}} \equiv [\frac{\partial \underline{u}}{\partial a}, \frac{\partial \underline{u}}{\partial b}, \frac{\partial \underline{u}}{\partial \gamma}]$ and

$$\hat{\sigma}_u^2 = \frac{1}{T} \sum_{t=1}^T (d_t - \hat{a} + \hat{b} p_t - \frac{\hat{b}}{\hat{\gamma}} p_{t-1})^2.$$

The restriction that $a=0$ and $b=1$ is tested by computing the following statistic:

$$\chi^2_{(2)} = (\hat{\underline{a}}_{b=1})' V^{-1} (\hat{\underline{a}}_{b=1}),$$

where V is the variance-covariance matrix for the estimates \hat{a} and \hat{b} . The test statistic has a large sample distribution that is Chi-squared with two degrees of freedom.

To estimate the parameters of equation (4), we must use Hansen's GMM estimator. To develop this estimator, first rewrite equation (4) as a function for u_t :

$$u_t = \frac{D_t}{D_{t-1}} - a + b \frac{P_t}{D_{t-1}} - \frac{b}{\beta} \frac{P_{t-1}}{D_{t-1}},$$

and let \underline{u} be a $T \times 1$ vector containing u_t , $t=1, \dots, T$. The instrumental variables for this estimator are included in a vector \underline{z}_t :

$$\underline{z}_t' = \left(1, \frac{D_{t-1}}{D_{t-2}}, \frac{P_{t-1}}{D_{t-1}}, \frac{X_{t-1}}{X_{t-2}}, \frac{P_{t-1}}{X_{t-1}}\right),$$

where X_t is earnings. Now form a matrix Z for the observations on the instruments so that \underline{z}_t' is the t 'th row of Z , a $T \times 5$ matrix. Let $\underline{\theta}' = (a, b, \beta)$ and form the GMM estimator as follows:

$$\min_{\underline{\theta}} \ell = \frac{1}{T} \underline{u}' Z W_T^{-1} Z' \underline{u},$$

where W_T is a weighting matrix which can be a function of sample data. Hansen shows that the optimal GMM estimator is the one in which W_T is equal to 2π times the spectral density matrix for $\underline{z}_t u_t$ evaluated at the zero frequency. In this paper, the matrix will be designated as S . In most cases, we must estimate this matrix. The following two-step estimator which replaces S with a consistent estimate \hat{S} is asymptotically equivalent to the optimal estimator: first, estimate $\underline{\theta}$ using a weighting matrix which produces initial consistent estimates; then use the consistent estimates $\hat{\underline{\theta}}$ to form u_t and estimate the spectrum of $\underline{z}_t u_t$ to

produce a consistent estimate \hat{S}_T ; finally, estimate $\underline{\theta}$ using \hat{S}_T as the weighting matrix. For the initial weighting matrix, I use $W_T = \frac{1}{T} Z'Z$, which yields the nonlinear two-stage least squares estimates which are consistent for $\underline{\theta}$. The smoothed periodogram and cross-periodogram estimators with flat windows are used to estimate S .⁷ The second step of the estimator is

$$\min_{\underline{\theta}} \ell = \frac{1}{T} \underline{u}'Z \hat{S}_T^{-1} Z' \underline{u}.$$

The covariance matrix for $\hat{\underline{\theta}}$ is

$$\text{Var}(\hat{\underline{\theta}}) = T \left[\left(\frac{\partial \underline{u}}{\partial \underline{\theta}} \right)' Z \hat{S}_T^{-1} Z' \left(\frac{\partial \underline{u}}{\partial \underline{\theta}} \right) \right]^{-1}.$$

The restriction that $a=0$ and $b=1$ is tested by computing a χ^2 statistic analogous to the previous one.

IV. Empirical Results

The present value model of stock prices is tested by using quarterly data for the Standard and Poor's Index of 500 Common Stocks for the period 1947 to the second quarter of 1983. In its publication Trade and Security Statistics, Standard and Poor's compiles quarterly data on its price index of 500 common stocks as well as indices on earnings and dividends for the companies included in the index.⁸ The implicit price deflator for personal consumption expenditures is used to deflate the three series. The estimates for equations (3) and (4) are presented in Tables I and II, respectively. The equations have been estimated for two sample periods: the first sample period contains the longer period of 1947:I to 1983:II and the second sample period contains only the more recent period of 1960:I to 1983:II.

We must detrend the three series prior to estimating equation (3), and I have used a method similar to that used by Shiller. The following regressions are used to estimate the common growth trend for real dividends and real stock prices:

$$\ln D_t = \hat{a}_1 + \ln(\hat{1+g}_1) \cdot t,$$

$$\ln P_t = \hat{a}_2 + \ln(\hat{1+g}_2) \cdot t.$$

The two estimates, $\ln(\hat{1+g}_1)$ and $\ln(\hat{1+g}_2)$, are averaged to form the OLS estimate of the common growth rate $(1+g)$, and this estimate is used to detrend both series. A separate trend regression is estimated for real earnings. For the longer sample period, the trend coefficient for $\ln D_t$ is .004398 and for $\ln P_t$ it is .005628. The resulting estimate of $(1+g)$ is 1.005026. The results for the longer sample period are contained in the first column of Table I. Using the z statistics for a and b, one can reject the null hypotheses that $a=0$ and $b=1$ at standard significance levels: the z statistic for a is 12.18 and for the test that $b=1$ it is -91.69. The $\chi^2_{(2)}$ test statistic for the joint hypothesis has an extreme value of 12405, so that the marginal significance is effectively zero. For the more recent sample period, the estimates change, but the hypothesis tests lead to the same conclusion. The trend coefficient for real dividends is lower, and it is negative for stock prices. This result alone indicates that the data do not fit the model, or that stock prices deviated substantially from their long-term trend. The results for the more recent sample period are contained in the second column of Table I. The z statistic for a is 11.92 and the z statistic for the test that $b=1$

is -159.95. The $\chi^2_{(2)}$ test statistic has a value of 44539 so that the marginal significance is still effectively zero. The Durbin-Watson statistic is also quite low suggesting that the model is inadequate. These results are consistent with the previous results of the variance bounds tests, and it is clear that the test based on the IV estimator is much more powerful.

The results for equation (4) are equally damaging. For this equation, it is not necessary to detrend the data; the effects of long-term trends are removed when we form the ratios $(\frac{D_t}{D_{t-1}}, \frac{P_t}{D_{t-1}}, \frac{P_{t-1}}{D_{t-1}})$. The estimates for the longer period are presented in the first column and those for the more recent period are in the second column. The $\chi^2_{(2)}$ statistic for the longer period is 5888 and for the more recent period it is 246356. The marginal significance level in both cases is effectively zero. The D.W. statistics support my concern that the error term in equation (4) is serially correlated. The tests for equation (4) indicate that the present value model is also rejected by the data if we model percentage changes in dividends and stock prices as stationary processes. These results do not substantiate this criticism of the previous work based on variance bounds tests.

V. Conclusion

The results for the S&P 500 data suggest that the present value model of stock prices with rational expectations does not adequately describe the behavior of asset prices in the stock market. The alternative test described here does not require some of the additional assumptions required to derive tests based on direct estimates of

variances, and the results are much more conclusive than those previously obtained for tests of variance bounds. In addition, this test is modified to handle the case in which the percentage changes in real dividends and real stock prices are stationary. The null hypothesis that has been rejected is a joint hypothesis which includes the present value model of stock prices and rational expectations. Rejection implies that at least one part of the joint hypothesis has been rejected by the data. These tests cannot distinguish which part of the joint hypothesis has been rejected, and it is possible that both parts have been rejected. One alternative hypothesis is that expected real rates of return and real interest rates are not constant. Indeed, the results suggest that much of the variation in stock prices is due to something other than variation in dividends. This view of the stock market is consistent with observations that stock prices change when interest rates change even though there may be no new information to change expectations about future dividends and earnings.

FOOTNOTES

1. For a discussion of this model and other expectations models, see LeRoy's survey (1982).
2. In most cases we must estimate the value of the discount factor and I have shown elsewhere (1982, pp. 114, 190-93) that the frequency domain variance estimator with a consistent estimate replacing the value of the discount factor is not asymptotically equivalent to the variance estimator which uses the true value of the discount factor.
3. For the asymptotic distribution of sample variances for time series, see Hannan (1970, pp. 209-12). In a footnote (footnote 15, p. 567), LeRoy and Porter recognize this subtle assumption in the derivation of their tests.
4. This subtle point is frequently ignored in applied econometrics. To apply Hansen's GMM estimator for example, we need to have stationary time series.
5. In a comment on Shiller's paper, Copeland (1983) has argued that the long term growth rate for dividends may be changing. His argument is confined to the effect of a one-period change, but we must consider how the growth rate changes over time. When we assume that the percentage changes in dividend and stock prices are stationary, we normally think of a constant growth rate for the series. If the random growth rate for dividends is itself a stationary process, then the series $(\frac{\Delta D}{D}, \frac{\Delta P}{P})$ will be stationary. If we assume that the random growth rate is not stationary, then we again run the risk of not satisfying the terminal condition for the present value model.
6. This estimator is the nonlinear two-stage least squares estimator described in Amemiya (1974), which is also a special case of Hansen's GMM estimator. This nonlinear minimization problem can be easily solved by replacing b/γ with a new parameter, say h , and then solving the linear first order conditions for \hat{a} , \hat{b} , and h . $\hat{\gamma}$ is then computed as $\hat{\gamma} = \hat{b}/\hat{h}$.
7. For consistent estimators of the spectrum, see Hannan (1970, pp. 273-88), Nerlove, Grether, and Carvalho (1979, pp. 57-68), and Koopmans (1974, Ch. 8). Some sources call the estimator used here the Daniell estimator. I also used the Parzen procedure described in Nerlove, Grether, and Carvalho (p. 67). Specifically, fourth-order autoregressions were estimated for each series and the estimated filters were used to prewhiten the data. The spectral estimates were then recolored by using the transfer functions associated with the filters.
8. The data for dividends are reported as a twelve-month moving total. LeRoy and Porter obtained the data necessary to recompute the quarterly dividends and their numbers are available in a technical appendix to their original paper.

TABLE I

RESULTS FOR ESTIMATION OF EQUATION (3)

$$d_t = a - b p_t + \frac{b}{\gamma} p_{t-1} + u_t$$

	Sample Period <u>1947:I to 1983:II</u>	Sample Period <u>1960:I to 1983:II</u>
\hat{a}	.4059	.9775
Standard error	(.03332)	(.08202)
\hat{b}	.01741	.007920
Standard error	(.01072)	(.006203)
$\hat{\gamma}$.8998	1.092
Standard error	(.07415)	(.09944)
$\chi^2_{(2)}$	12405.	44539.
$\hat{\sigma}_u^2$.01276	.01488
T	145	93
D.W.	2.06	.81
<u>Trend Coefficients</u>		
Dividends	.004398	.001158
Stock Price	.005628	-.005518
Earnings	.005572	.005267
$(1+\hat{g})$	1.005026	.9978225

TABLE II

RESULTS FOR ESTIMATION OF EQUATION (4)

$$\frac{D_t}{D_{t-1}} = a - b \frac{P_t}{D_{t-1}} + \frac{b}{\beta} \frac{P_{t-1}}{D_{t-1}} + u_t$$

	Sample Period <u>1947:I to 1983:II</u>	Sample Period <u>1960:I to 1983:II</u>
\hat{a}	.8095	.9492
Standard error	(.08783)	(.03752)
\hat{b}	.02249	.004810
Standard error	(.01548)	(.002007)
$\hat{\beta}$.9089	.9033
Standard error	(.07598)	(.06250)
$\chi^2_{(2)}$	5888.	246356.
T	144	92
D.W.	3.50	2.85

APPENDIX

To examine the properties of the error term in equation (3), we need to analyze the time series representation for d_t and p_t and the restrictions implied under the alternative hypothesis to the present value model. Let d_t have the following form in a linear multiple time series representation that is similar to equation (3) in LeRoy and Porter:

$$d_t = c + \underline{\delta}'_0 \underline{\varepsilon}_t + \underline{\delta}'_1 \underline{\varepsilon}_{t-1} + \dots = c + \underline{\delta}'(L) \underline{\varepsilon}_t.$$

Under the alternative hypothesis, we let market expectations be formed as follows:

$$\hat{E}_t(d_{t+j} - c) = \frac{1}{b} E_t(d_{t+j} - c),$$

where \hat{E}_t is the market's conditional expectation. The market systematically overestimates the deviation of dividends from its trend if $b < 1$ and underestimates the deviation if $b > 1$. If b equals one, we have the present value model. The model now requires p_t to have the following time series representation:

$$\begin{aligned} p_t &= \hat{E}_t \sum_{j=1}^{\infty} \gamma^j d_{t+j} = \frac{c\gamma(b-1)}{(1-\gamma)b} + \frac{1}{b} E_t \sum_{j=1}^{\infty} \gamma^j d_{t+j} \\ &= \frac{c\gamma}{1-\gamma} + \frac{1}{b} [\underline{a}'_0 \underline{\varepsilon}_t + \underline{a}'_1 \underline{\varepsilon}_{t-1} + \dots] \\ &= \frac{c\gamma}{1-\gamma} + \frac{1}{b} \underline{a}'(L) \underline{\varepsilon}_t, \end{aligned}$$

where $\underline{a}'_j = \sum_{k=1}^{\infty} \gamma^k \underline{\delta}'_{k+j}$. We now examine the error term u_t . The intercept a in equation (3) equals $c(1-b)$.

$$\begin{aligned} u_t &= d_t - a + b p_t - \frac{b}{\gamma} p_{t-1} \\ &= c + \underline{\delta}'(L) \underline{\varepsilon}_t - c(1-b) + \frac{c\gamma b}{1-\gamma} + \underline{a}'(L) \underline{\varepsilon}_t - \frac{cb}{1-\gamma} - \frac{1}{\gamma} \underline{a}'(L) \underline{\varepsilon}_{t-1}. \end{aligned}$$

It can be easily shown that all the constants on the right hand side of the equation cancel. Next, we match the coefficients on $\underline{\varepsilon}_{t-1}$, $\underline{\varepsilon}_{t-2}$, ... and show

$$\begin{aligned} \underline{\delta}'_j \underline{\varepsilon}_{t-j} + \underline{a}'_j \underline{\varepsilon}_{t-j} - \frac{1}{\gamma} \underline{a}'_{j-1} \underline{\varepsilon}_{t-j} &= \left(\underline{\delta}'_j + \sum_{k=1}^{\infty} \gamma^k \underline{\delta}'_{k+j} - \frac{1}{\gamma} \sum_{k=1}^{\infty} \gamma^k \underline{\delta}'_{k+j-1} \right) \underline{\varepsilon}_{t-j} \\ &= 0 \quad \text{for } j = 1, 2, 3, \dots \end{aligned}$$

Hence, the coefficients on $\underline{\varepsilon}_{t-1}$, $\underline{\varepsilon}_{t-2}$, are all zero and we have for the error term:

$$u_t = \left(\underline{a}'_0 + \underline{\delta}'_0 \right) \underline{\varepsilon}_t = \left(\sum_{k=0}^{\infty} \gamma^k \underline{\delta}'_k \right) \underline{\varepsilon}_t.$$

The error term is a linear combination of only the current innovations of the dividend series and is therefore serially uncorrelated. If we let $b=1$, then the analysis also applies for the present value model.

REFERENCES

- Takeshi Amemiya, "The Nonlinear Two-Stage Least-Square Estimator," Journal of Econometrics 2 (1974): 105-10.
- Basil L. Copeland, "Do Stock Prices Move Too Much to be Justified by Subsequent Changes in Dividends?: Comment," American Economic Review (March 1983): 234-35.
- Eugene F. Fama and Harvey Babiak, "Dividend Policy: An Empirical Analysis," Journal of the American Statistical Association (December 1968): 1132-61.
- Marjorie Flavin, "Small Sample Bias in Tests of Excess Volatility in the Bond Market," University of Virginia, January 1982.
- John Geweke, "A Note on the Testable Implications of Expectations Models," University of Wisconsin-Madison, December 1979.
- Clive W. J. Granger, "Some Consequences of the Valuation Model When Expectations are Taken to be Optimum Forecasts," Journal of Finance (March 1975): 135-45.
- E. J. Hannan, Multiple Time Series (New York: John Wiley & Sons, 1970).
- Lars P. Hansen, "Large Sample Properties of Generalized Method of Moments Estimators," Econometrica (July 1982): 1029-54.
- L. H. Koopmans, The Spectral Analysis of Time Series (New York: Academic Press, 1974).
- Stephen F. LeRoy, "Expectations Models of Asset Prices: A Survey of Theory," Journal of Finance (March 1982): 185-217.
- _____ and Richard D. Porter, "The Present-Value Relation: Tests Based on Implied Variance Bounds," Econometrica (May 1981): 555-74.
- John Lintner, "Distribution of Incomes of Corporations Among Dividends, Retained Earnings, and Taxes," American Economic Review (May 1956): 97-113.
- Marc Nerlove, David M. Grether, and Jose L. Carvalho, Analysis of Economic Time Series: A Synthesis (New York: Academic Press, 1979).
- Louis O. Scott, "Risk Aversion, Rational Expectations, and Constant Expected Real Rates of Return," Unpublished Ph.D. dissertation, University of Virginia, 1982.
- Robert J. Shiller, "The Use of Volatility Measures in Assessing Market Efficiency," Journal of Finance (May 1981a): 291-304.

_____, "Do Stock Prices Move Too Much to be Justified by Subsequent Changes in Dividends?" American Economic Review (June 1981b): 421-36.

_____, "Reply," American Economic Review (March 1983): 236-37.

Kenneth J. Singleton, "Expectation Models of the Term Structure and Implied Variance Bounds," Journal of Political Economy (December 1980): 1159-76.

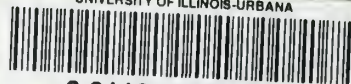
HECKMAN
BINDERY INC.



JUN 95

ound-To-Place® N. MANCHESTER,
INDIANA 46962

UNIVERSITY OF ILLINOIS-URBANA



3 0112 060296123